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ORIGINAL PAPER

Education and gender role attitudes

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Abstract
This paper examines whether education plays an important role in shaping individuals' gender role attitudes. We exploit exogenous variation in temporal and geographical impacts of the 1986 Compulsory Education Law in China, which reduced inequality in compulsory school attendance across regions. Using the data from the China General Social Survey, we find that the extra schooling induced by the compulsory schooling reform leads to more egalitarian gender role attitudes. Education's liberalizing effect is concentrated among females and urban residents. However, education's impacts on gender-equal behavior are much weaker than impacts on attitudes. Finally, we discuss the potential channels through which education shapes individuals' gender-role attitudes.

Keywords Education · Gender role attitudes · Compulsory education law · China

JEL Classification I20 · J16

1 Introduction

China has experienced tremendous changes in cultural beliefs about the appropriate or natural role of a woman in society, which we refer to as gender role attitudes. A definite

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trend toward more egalitarian conceptions of women's roles is observed in China. However, despite the substantial improvement of women's socioeconomic status, some traditional gender norms persist. Reforming basic gender norms may potentially produce long-lasting improvements in outcomes for women and society (Duflo 2012; Zhang 2019); thus, understanding the factors that can shape individuals' gender role attitudes has important policy implications. This paper examines whether education plays a role in shaping individuals' gender role attitudes and influencing their gender-equal behavior. To this end, we exploit exogenous temporal and geographical variation in the enforcement of the 1986 Compulsory Schooling Laws (CSLs) in China.

The goals of education include more than mere cognitive and intellectual development of individuals. Education has also been a major institution shaping individuals' attitudes, values, and beliefs.[1] In particular, education promotes a more enlightened world outlook.[2] Previous studies have shown that individuals with higher educational attainment tend to hold more egalitarian attitudes toward gender roles than those with less educational attainment, suggesting that receiving more education may change individuals' attitudes (Thornton et al. 1983a, b; Kane 1995; Brewster and Padavic 2000). Unfortunately, most of the existing studies indicate only a correlation between education and egalitarian gender role attitudes, and very few credibly investigate causal links.[3] The key difficulty in estimating the causal effect of education on attitudes is that some unobserved characteristics affecting gender role attitudes are likely to influence schooling decisions. For example, mothers' gender role attitudes and experiences, which play an important role in shaping their children's attitudes, may also affect their children's educational attainment.

This paper contributes to the literature by identifying the causal effects of education on gender role attitudes. We exploit exogenous temporal and geographical variation in the enforcement of CSLs and cross-province variation in pre-reform educational levels to estimate the causal effect of education. Although China's nationwide Compulsory Schooling Law officially went into effect on 1 July 1986, provinces/municipalities were allowed to have different effective dates for implementing the CSLs. Cohorts that were above 15 years old when the CSLs took effect were not obliged to complete 9-year (a 6-year primary school and a 3-year junior high school) compulsory schooling. In contrast, later cohorts experienced an exogenous increase in educational attainment. Moreover, provinces differ remarkably in average compulsory school attendance before the enforcement of the CSLs. Because the law requires everyone to complete at least junior high school, provinces with lower junior high school completion rates gained more from the implementation of the CSLs, while provinces with higher completion rates gained little and thus can serve as a control group. With the temporal and geographic variation in the effective dates of the CSLs and the cross-province

[1] A number of studies have shown that education reduces religiosity, religious acts, and superstitious belief (Hungerman 2014; Mocan and Pogorelova 2017), and higher education leads to a more positive reported attitude toward immigrants (d'Hombres and Nunziata 2016).

[2] For instance, education promotes democratic norms of equality and civil rights and fosters egalitarian racial attitudes (Hyman and Wright 1979).

[3] Dinçer et al. (2014) use Turkey's 1997 Education Law to study the effect of education on women's attitudes toward gender equality and find little evidence that schooling changed women's attitudes toward gender equality. Erten and Keskin (2018) also examine the effect of education on gender role attitudes using the same educational reform. They find a significant effect of education on only two gender role attitudes: men should also do housework, and men in the family are responsible for a woman's behavior.

variations in pre-reform educational levels, we apply a difference-in-difference-in-difference (triple-difference) framework to investigate the impact of the CSLs on schooling. People who were born later and in an area with lower pre-reform educational levels were more exposed to the CSLs. Therefore, we compare cohorts based on their year of birth relative to the implementation of the CSLs in their province of birth and the educational level before the enforcement of the CSLs in their province of birth.

Using the data from the China General Social Survey (CGSS), we find that changes in the CSLs have a significant effect on educational attainment. Using the instrumental variable approach, we show that the extra schooling induced by the 1986 Compulsory Schooling Laws in China leads to more egalitarian gender role attitudes. Education's liberalizing effect is markedly stronger for females and urban residents. Education shapes individuals' gender role attitudes and also influences the participation of women in activities performed outside home. Moreover, women with higher educational attainment have a lower probability of giving birth to a second child when their first birth is a girl. However, women's education does not significantly affect the relative division of household chores between men and women in a household. The results provide suggestive evidence that the impacts of education on behavior are much weaker than impacts on attitudes.

In recent years, economists have shown increasing interest in cultural norms and beliefs. A vast body of literature has documented the role of norms and beliefs in explaining economic outcomes. In particular, norms or beliefs about the appropriate role of women in society affect females' participation in paid and unpaid work. For example, traditional gender role attitudes imply more home production of goods and services and less participation in market activities for women (Alesina and Giuliano 2010). Egalitarian views toward gender roles and work values are positively associated with female labor force participation rates (Fortin 2005, 2015; Fernandez 2007; Chen and Ge 2018; Ye and Zhao 2018).[4] The preferences and beliefs regarding women's roles in the countries of ancestry are economically and statistically significant in explaining women's fertility decisions (Kaufman 2000; Fernandez and Fogli 2006, 2009). Gender norms also shape the distribution of unpaid work within a household (Juster and Stafford 1991; Alvarez and Miles 2003; Hwang et al. 2017). Moreover, a large literature, surveyed by Jayachandran (2015), studies the role of social norms about gender roles in explaining gender inequality in developing countries.

Although most studies on social norms examine the impact of cultural norms, in particular, gender role attitudes, on social and economic outcomes, few studies focus on the formation of these attitudes.[5] One line of research emphasizes the deep historical roots of gender norms (see Giuliano 2017, for a review). For example, gender role attitudes are found to be associated with religiosity (Seguino 2011), agricultural history (Hansen et al. 2015), and the agricultural environment faced by the first settled farming communities (Alesina et al. 2013). Another line of research focuses on short-term determinants of gender role attitudes. Many believe that gender role attitudes are largely determined early in childhood and family is the primary influence on gender role

[4] See Bertrand (2011) for an extensive survey on the relationship between social and gender identity norms and women's labor market choices and outcomes.

[5] A few studies have examined the factors that affect the formation of other beliefs and preferences than gender role attitudes. For example, Giuliano and Spilimbergo (2014) find that the experience of macroeconomic shocks when young has long-lasting effects on political and economic beliefs.

development during the early years of life. A large number of studies have documented intergenerational transmission of attitudes toward gender roles (Cunningham 2001; Fernandez et al. 2004; Fernandez and Fogli 2006; Farre and Vella 2013; Hwang 2016; Chen and Ge 2018; Dhar et al. 2019). Parents' attitudes and behavior, for example, parents' housework allocation, and mothers' gender role attitudes and labor force participation influence the offspring's attitudes toward gender roles and economic behavior. Other factors regarding family, for instance, child gender (Washington 2008) or sibling's gender (Healy and Malhotra 2013; Brenøe 2018), also strongly impact individuals' gender role attitudes. Additionally, the introduction of cable television, which exposes viewers to new information about the outside world, affects the gender attitudes of individuals in rural India (Jensen and Oster 2009). The exposure to a female leader also weakens stereotypes about gender roles in the public and domestic spheres (Beaman et al. 2009).

Our findings contribute to a deeper understanding of the formation of gender role attitudes. Gender inequality is widely observed across countries of all income levels but is especially severe in developing countries. Recent work has shown that cultural norms about the appropriate role of women in society contribute to gender inequality. Moreover, several studies have argued that social norms about gender roles might be amenable to change. For instance, the introduction of cable television (Jensen and Oster 2009), exposure to a female leader (Beaman et al. 2009), and exposure of men to women in a traditionally male-dominated environment (Dahl et al. 2018) lead to more progressive gender attitudes. This paper adds to this line of literature by providing additional evidence that education has shaped the evolution of norms and beliefs about the appropriate role of women in society. We also discuss potential channels that education's liberalizing effect may operate through, including providing a gender-neutral curriculum, increasing the economic value of women, and changing the way people get access to new information outside.

This study is also related to the strand of literature showing that education offers a wide range of benefits that extend beyond increases in labor market productivity.[6] For example, a large number of studies estimate the causal impact of education on health, measured by adult mortality (Lleras-Muney 2005, 2006) or self-reported health status (Huang 2015). Milligan et al. (2004) find that schooling improves civic participation in the United States. Lochner and Moretti (2004) show that schooling reduces the probability of incarceration and arrest. Heckman et al. (2018) analyze a range of non-market benefits of education and ability, including incarceration, mental health, voter participation, trust, and participation in welfare. Several studies document the causal effect of education on the formation of (dis)religious beliefs (Brown and Taylor 2007; Hungerman 2014; Mocan and Pogorelova 2017). Our study contributes to the literature by demonstrating that education can promote a more enlightened world outlook, i.e., more egalitarian attitudes toward gender roles.

The remainder of the paper is organized as follows. In Sect. 2, we provide a brief overview of the evolution of gender role attitudes in China. Section 3 describes the dataset used in the analysis. Section 4 outlines the empirical specification. We present the results in Sect. 5. Section 6 concludes.

[6] See Lochner (2011) for a comprehensive review of the literature.

2 Gender role attitudes in China

The traditional Chinese family was characterized as patriarchal and patrilineal. Women were at a severe social disadvantage relative to men (Thornton and Lin 1994). From the Han dynasty onward in imperial China, Confucianism largely defined the mainstream discourse on gender. The *Three Obediences* and *Four Virtues* were a set of basic moral principles specifically for women in Confucianism. According to these principles, a virtuous woman must follow the lead of the males in her family. That is, she should obey her father before her marriage, obey her husband in marriage, and obey her son in widowhood. The four feminine virtues that women were expected to possess include morality, proper speech, modest manner and appearance, and diligent work. This set of virtues represents traditional society and the conventional roles of women within that society. Moreover, women were denied the opportunity for education, and their activities were confined to the domestic arena (Zhang 2019).[7] The long history of endorsement of Confucianism reinforced the obligatory gender roles and the notion of women's inferiority.

Since the collapse of the Qing dynasty in 1911, norms concerning gender roles have changed remarkably. The New Culture Movement of the mid-1910s and 1920s revolted against norms of the traditional Confucian culture, called for an end to the patriarchal family, and advocated individual freedom and women's liberation. Consequently, some women began to acquire formal education and work outside the home in order to fight for their economic independence.

After the founding of the People's Republic of China, China instituted the National Marriage Law in 1950, which formally legalized free choice in marriage and explicitly equalized the rights of husbands and wives in the family. This law also granted wives the freedom to participate in the labor market. In the same year, national law in China granted women legal rights to land for the first time. Women were granted the right to vote even earlier, in 1947. To promote gender equality, the Communist party advanced the slogan, "women hold up half the sky" in 1968 to illustrate the importance of women's contributions to the society. Women's socioeconomic status has been significantly improved. This is most significantly shown in increasing employment opportunities and educational attainment of women. For instance, in 1982, the percentages receiving postsecondary education were respectively 1.45% for men (aged between 16 and 60) and 0.55% for women; these percentages increased to 9.00 and 7.33% in 2005. The labor force participation rate was 93% for women between the prime ages of 16 and 55 in 1988 (Ge and Yang 2014).

Despite these substantial improvements, some of the traditional norms concerning gender identity still exist. For example, Yu and Yu (2012) show large gender gaps in household work allocation in contemporary China, with the lion's share falling on the shoulders of the wife rather than the husband. Ge and Yang (2014) find that the rate of female labor force participation started to decline in the 1980s, along with a widening gender gap in earnings. Therefore, it is of great empirical importance to investigate whether education plays an important role in shaping norms and beliefs regarding gender roles.

[7] The Chinese proverb "A woman's lack of talent is in itself a virtue" describes the ideal wife candidate for Chinese men in a traditional patriarchal society.

3 Data

In this paper, we use the data from the China General Social Survey (CGSS), which has been collected since 2003 jointly by the Hong Kong University of Science and Technology and Renmin University of China. The CGSS is designed to be the Chinese counterpart of the General Social Survey (GSS) in the USA. The CGSS is a multistage stratified probability proportional to size (PPS) sampling survey of the Chinese population in mainland China and is designed to be representative for the Chinese population after weighting. The dataset contains rich information on the demographic characteristics, social attitudes, and labor market outcomes of each respondent. It is a repeated cross-sectional data set. For the purpose of our paper, we use the four waves of the survey conducted in 2010, 2012, 2013, and 2015.[8] In this paper, we restrict our analysis to individuals born between 1960 and 1990. This leaves us with a sample of 25,009 individuals over the sample period.

The 2010, 2012, 2013, and 2015 waves of CGSS have several attitudinal questions regarding the respondents' opinions about gender roles.[9] More specifically, the respondents were asked whether they "totally disagree," "somewhat disagree," "are neutral," "somewhat agree," or "totally agree" with each of the following five statements:

(1) Men should focus on career, whereas women should focus on family (*career*);
(2) Men are naturally more competent than women (*competency*);
(3) Getting married to a good man is more important than having a good job (*marriage*);
(4) Female workers should be dismissed first during a recession (*dismissal*);
(5) Husbands and wives should equally share household chores (*housework*).

These questions reflect respondents' attitudes about the gender role specialization and woman's rights to a career. We recode the responses to the first four statements into a binary variable by combining "totally disagree" and "somewhat disagree" as $\text{Attitude}_i = 1$, which represents the egalitarian or modern gender role attitude; and combining "neutral," "somewhat agree," and "totally agree" as $\text{Attitude}_i = 0$, which represents the traditional gender role attitude. For the last statement, we combine "totally agree," "somewhat agree," and "neutral" as $\text{Attitude}_i = 1$ and "totally disagree" and "somewhat disagree" as $\text{Attitude}_i = 0$.

We also construct an index to measure the overall attitudes toward gender roles. To this end, we first sum up the responses (Attitude_i) to the above five statements, ranging from 0 to 5. To facilitate interpretation, we normalize the index so that it has a mean of

[8] The 2010 wave of CGSS covers all 31 provinces/municipalities in mainland China and 11,783 households. The 2012 CGSS covers 29 provinces/municipalities (excluding Hainan and Tibet) and 11,765 households; the 2013 and 2015 CGSS cover 28 provinces/municipalities (excluding Hainan, Tibet, and Xinjiang) and 11,438 and 10,968 households, respectively.

[9] Our estimates of schooling effects may be biased if the likelihood of not answering the questions concerning gender-role attitudes is affected by educational attainment. However, the problem is not severe in our sample since only 0.14–0.88% of respondents did not report their gender role attitudes.

zero and a standard deviation of one. The summary statistics of gender-role attitudes and other variables for the CGSS sample are presented in Appendix 3. On average, people hold the least gender-equal attitudes toward career and the most gender-equal attitudes toward the division of housework.

Table 1 presents differences in gender role attitudes across education groups. Overall, we find that individuals with more schooling are more likely to hold egalitarian gender role attitudes. Consistent with previous findings, the odds of holding egalitarian attitudes are monotonically increasing with education for all measures of attitudes and for both men and women (Thornton et al. 1983a, b; Kane 1995; Brewster and Padavic 2000). Additionally, egalitarian gender role attitudes are more prevalent among educated women than educated men.

Table 1 Gender role attitudes by education level

Attitude	Education	Full sample	Male	Female
Career	Illiterate	11.99	10.84	12.31
	Primary school	14.96	14.92	14.99
	Junior high school	25.39	23.09	27.69
	Senior high school	34.7	28.6	41.49
	College and above	42.88	34.44	51.99
Competency	Illiterate	25.26	28.07	24.5
	Primary school	31.47	32.44	30.81
	Junior high school	44.68	42.38	46.98
	Senior high school	51.98	44.93	59.85
	College and above	58.57	51.02	66.71
Marriage	Illiterate	21.28	24.47	20.42
	Primary school	26.93	29.91	24.88
	Junior high school	36.47	38.28	34.65
	Senior high school	39.71	38.14	41.46
	College and above	43.44	39.74	47.43
Dismissal	Illiterate	58.72	54.48	59.88
	Primary school	64.22	60.47	66.81
	Junior high school	73.73	69.67	77.8
	Senior high school	76.81	70.57	83.75
	College and above	81.77	76.17	87.8
Housework	Illiterate	84.02	79.3	85.31
	Primary school	84.47	81.76	86.33
	Junior middle school	85.93	83.35	88.53
	Senior high school	87.21	83.37	91.49
	College and above	88.75	85.2	92.57

Note: Authors' own calculation based on 2010, 2012, 2013, and 2015 waves of CGSS. Sample consists of individuals born between 1960 and 1990. This table presents the percentage of individuals holding egalitarian gender role attitudes.

4 Empirical strategy

4.1 Impact of education on gender role attitudes

To investigate the causal effects of education on gender role attitudes, we estimate the following equation:

$$\text{Attitude}_{\text{ipt}} = \delta_0 + \delta_1 \text{Edu}_{\text{ipt}} + \delta_2 X_{\text{ipt}} + D_{\text{ipt}} + u_{\text{ipt}}, \tag{1}$$

where $\text{Attitude}_{\text{ipt}}$ denotes reported gender role attitudes of individual i born in year t and receiving primary and junior high school education in province p. $\text{Attitude}_{\text{ipt}}$ equals to 1 if individual i holds egalitarian gender role attitudes, and zero otherwise. Edu_{ipt} is the years of schooling of individual *i*. δ_1 is the coefficient of interest, which captures the effect of one additional year of schooling on the endorsement of egalitarian gender role attitudes. X_{ipt} refers to a vector of control variables, including gender, age, ethnicity, household registration (*hukou*) status (urban/rural), both parents' years of schooling, and mother's employment status when the respondent was 14.[10] D_{ipt} denotes a set of dummies that account for survey year fixed effects, province fixed effects, and their interactions, birth year fixed effects, and birth month fixed effects.

In estimating Eq. (1), two important difficulties may arise. First, schooling may not be exogenous. Gender stereotypes affect individuals' educational choices (Favara 2012), which may induce a reverse causality problem. Moreover, mothers' gender role attitudes and experiences, which play an important role in shaping the attitudes of their offspring, may also affect the educational attainment of their children. The 1986 *Compulsory Education Law* in China provides an exogenous source of variation in educational attainment that we can use to identify the causal impact of schooling on gender role attitudes. A second problem that may arise in the estimation of Eq. (1) is the existence of reporting error, namely, gender role attitudes are self-reported. Better educated respondents may tend to know how to properly respond to the attitudinal questions and hide their "true" attitudes. Consequently, they are more likely to report egalitarian gender role attitudes. We address the reporting error problem by examining the impact of education on individuals' gender-equal behavior in Sect. 4.4.

4.2 Compulsory education Laws in China

China's nationwide Compulsory Education Law was passed on April 12, 1986 and officially went into effect on 1 July 1986. This was the first time that China used a law to specify educational policies for the entire country. This law had several important features (China Ministry of Education 1986). First, 9 years of education became compulsory. Second, all children who have reached the age of 6 shall enroll in school and receive their compulsory education. In this way, children were supposed to complete their compulsory education when they reached 15 years old. Those under

[10] The household registration (*hukou*) system in China, classifying each person as a rural or an urban resident, serves as an internal passport. The *Hukou* system is a major means of controlling internal migration and determining eligibility for state-provided services and welfare, such as education and medical care. *Hukou* refers to the registration of an individual in the system.

15 who had already left school by the law's effective date were required to return to school until they turned 15. Third, compulsory education was in principle free of charge.[11] Fourth, organizations and individuals were forbidden to employ school-age children or adolescents for work.[12] Fifth, local governments were allowed to collect education taxes to finance compulsory education (Fang et al. 2012).

The central government also planned to have different implement forces across different regions. Prior to the implementation of the law, the Chinese Communist Party (CCP) issued the *Decisions about the Education System Reform* in 1985, decentralizing educational responsibility to lower levels of governments. This reform increased the inequality of funding between rural and urban schools, leaving many local governments in poor provinces or regions with insufficient resources to fully implement the law. Thus, as documented in *Decisions about the Education System Reform*, the enforcement of the compulsory education law differed by three types of regions: (1) cities and economically developed areas in coastal provinces and a small number of developed areas in the hinterland; (2) towns and villages with medium-level development; and (3) economically backward areas. The central government sets different targets for each type of region. Moreover, the central government tried to support the less-developed regions.

As the central government recognized that not all provinces would have sufficient resources to enforce the law immediately, provinces were allowed to have different effective dates to implement the law. Figure 1 shows the effective dates of the CSLs in different provinces and municipalities in China. Most provinces and municipalities implemented the law in 1986 and 1987, while some provinces implemented the law after 1990, for example, Gansu, Guangxi, Hainan, and Tibet. Since then, China has made significant progress in the implementation of 9-year compulsory education. The country has basically achieved its goal of universal compulsory education by 2000, covering 85% of its population. As illustrated in Appendix 1, the net enrollment rate of primary-school-age children attained 99.2% in 2005. The promotion rate of primary school graduates to junior high school has increased from 69.5% in 1986 to about 87% in 1990 and reached 98.4% in 2005.

Importantly, regions differ remarkably in educational attainment prior to the CSLs. Figure 2a shows that the proportion of the population aged 16–18 with fewer than 9-year education in 1982 varies from 6.6% in Beijing to 90.5% in Tibet. Because the goal of the CSLs is to achieve universal compulsory education, low-education regions should experience more education increase then high-education regions. Figure 2b shows that the decline in the proportion of the population with fewer than 9-year schooling increases with the proportion prior to the CSLs. This cross-regional difference prior to the CSLs permits a treatment/control estimation strategy. Being born later relative to the effective year of the CSLs and in an area with lower pre-reform educational level implies more exposure to the benefits of the CSLs. The variation in the effective date of the reform across provinces and the variation in the pre-reform

[11] However, tuition-free 9-year compulsory education was not strictly enforced immediately after the enactment of the CSLs. The free compulsory education reform has been implemented in China since 2008 (Xiao et al. 2017).

[12] Employers who hired child laborers younger than age 16 would be penalized, including warnings, fines, suspension of business operations, and withdrawal of business licenses.

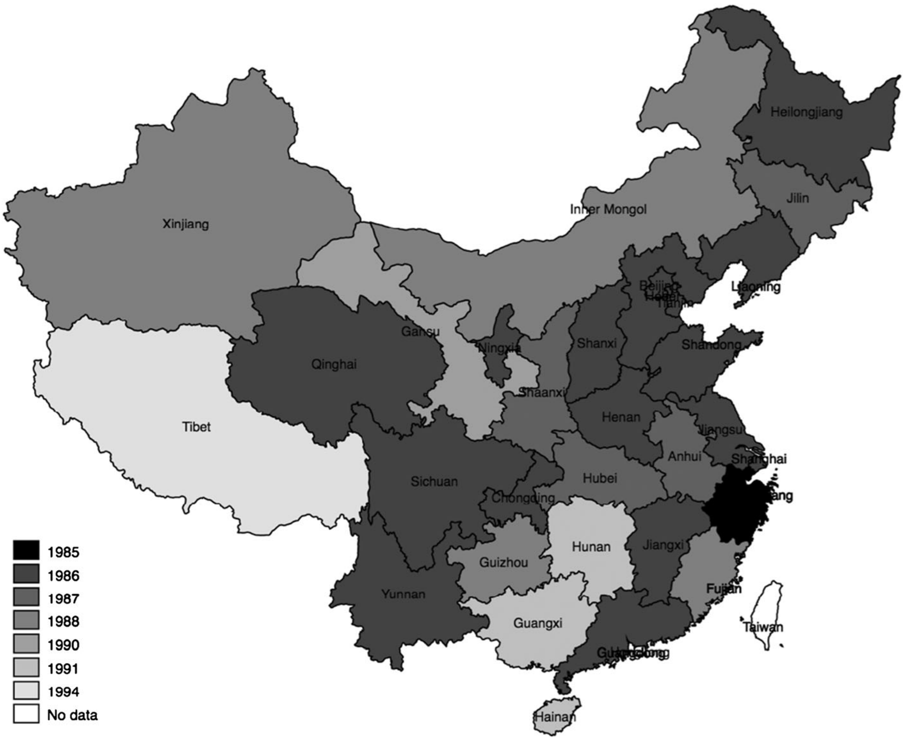

Fig. 1 The effective dates of the Compulsory Schooling Laws in different provinces. *Note*: Authors' own construction based on various local legal and administrative documents. Exact dates are listed in Appendix 4

educational level at the province level allows us to apply a triple-difference strategy to identify the causal impact of the CSLs on years of schooling in China.[13]

4.3 Impact of the CSLs on education and gender role attitudes

We use the exogenous temporal and geographical variation in the enforcement of the CSLs and across-province variation in pre-reform education, and estimate the impact of the CSLs on years of schooling using the following equation:

$$\mathrm{Edu}_{\mathrm{ipt}} = \beta_1 \mathrm{prop}_{\mathrm{p}}^{\mathrm{edu}<9} \times \mathrm{Exp}_{\mathrm{pt}} + \beta_2 X_{\mathrm{ipt}} + D_{\mathrm{ipt}} + D_{\mathrm{effective\ year}} \times D_{\mathrm{t}} + \sum_j x_{\mathrm{p}}^j \times \mathrm{Exp}_{\mathrm{pt}} + \varepsilon_{\mathrm{ipt}}, \tag{2}$$

[13] Several recent studies have explored the exogenous increase in education induced by the 1986 Compulsory Schooling Laws in China. Fang et al. (2012) and Xie and Mo (2014) apply a dichotomous measure of policy exposure to estimate the effects of education on income and health. Chen et al. (2018) employ Regression Probability Jump and Kink Design to identify the effect of education on non-cognitive skills. Cui et al. (2019) use within and between cohort variation in reform exposure to examine the impact of maternal education on the children. Huang (2015) exploits both the staggered adoption of the laws and the variation in pre-law educational levels across provinces to investigate the effects of education on health. The identification strategy used in our paper extends Huang (2015) by allowing the year of birth effects to vary across regions (Stephens and Yang 2014) and by ruling out other explanations of the differential changes in education across provinces (Bleakley 2010).

where Edu_{ipt} refers to years of schooling of individual i born in year t who received primary and junior high school education in province p. ε_{ipt} is the error term, clustered at the province-birth year level.[14]

Exp_{pt} is the CSLs reform exposure for birth cohort t in province p, which is one if the individual is fully eligible for the law (i.e., aged 6 or below when the law took effect) and zero if the individual is ineligible (i.e., aged 16 or above when the law took effect). For individuals aged between 6 and 15 when the law took effect, we assume that the CSLs reform exposure changes linearly with age (see Appendix 2).[15] The CSLs reform exposure is jointly determined by the effective year of the CSLs in the province where individuals received their compulsory education and their year of birth.

In this study, as the survey does not provide the information on the province where individuals received their compulsory education, we assume that the province of *hukou* is the province where an individual received his/her primary and junior high school education if the individual has been living in the *hukou* province till the survey time. If the province of current residence is different from the *hukou* province, we use the mother's province of residence when the child was born instead. One potential concern is that the mother's province of residence when the child was born may be different from the province where the child received his/her compulsory education. However, as the rural-urban migration was strictly restricted until the late 1980s in China, the interprovincial migration rate was low before the 1990s, indicating that the assumptions we make here may not be too strong.

$\text{prop}_p^{\text{edu}<9}$ denotes pre-reform educational level in province *p*, measured by the proportion of the population aged 16–18 who had received fewer than 9 years of schooling in province p in 1982.[16] The coefficient of interest is β_1, which captures the differential increase in education due to the CSLs between the provinces with lower and higher prior education. We expect that $\beta_1 > 0$, which implies a greater increase in years of schooling after the enforcement of CSLs for people in provinces with lower pre-reform education.

In addition to the fixed effects (D_{ipt}) that we control for in Eq. (1), we also include effective year—enrollment year fixed effects. The variable that captures the pure effect of the exposure to the CSLs is therefore absorbed by these fixed effects. We, therefore, exploit the fact that eligible individuals should comply more with the CSLs if they were from low-education regions than if they were from high-education regions. We also include the interactions between reform exposure and pre-reform characteristics of the province of birth, which include average wage, population density (person/km^2), highway mileage (kilometer), and fiscal expenditure on culture, education, science, and public health in 1985. Thus, we rule out other factors potentially causing differential changes in education across provinces and attribute the estimated CSLs impact to differential pre-reform educational level.

To estimate the schooling impacts on gender role attitudes, we can apply a reduced form regression specification, where education on the left-hand side of Eq. (2) is

[14] As a robustness check, we also cluster the error term at the province level. The results, which are available from the authors upon request, do not change much.

[15] The results do not rely on the linear-function assumption. The results applying a step function are largely unchanged and are available from the authors upon request.

[16] The proportion of the population aged 16–18 who had received fewer than 9 years of schooling prior to the CSLs is calculated based on the 1% sample of 1982 census data.

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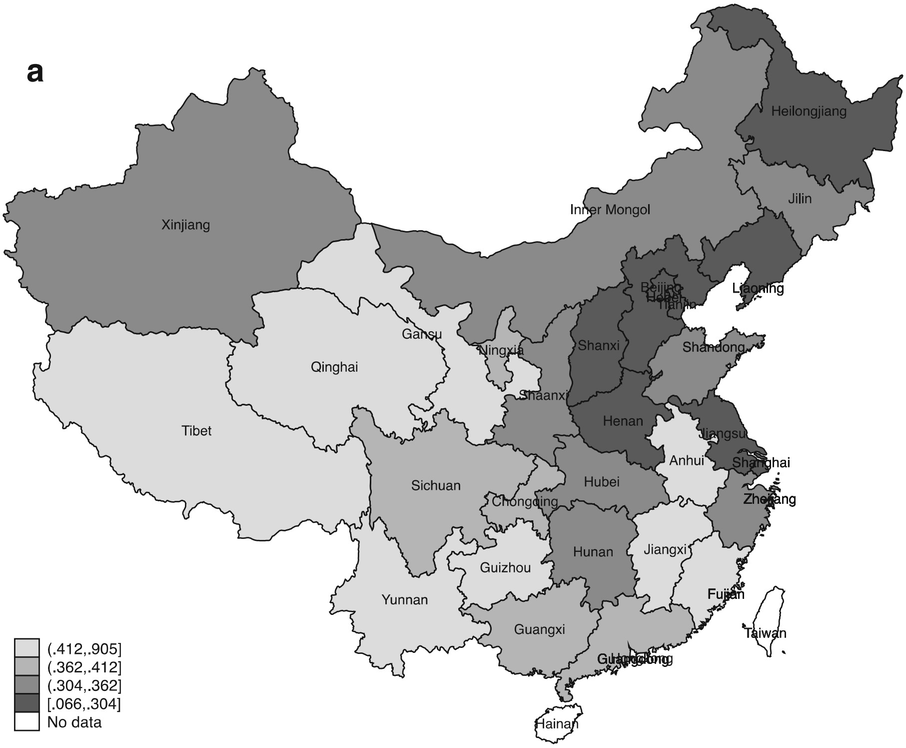
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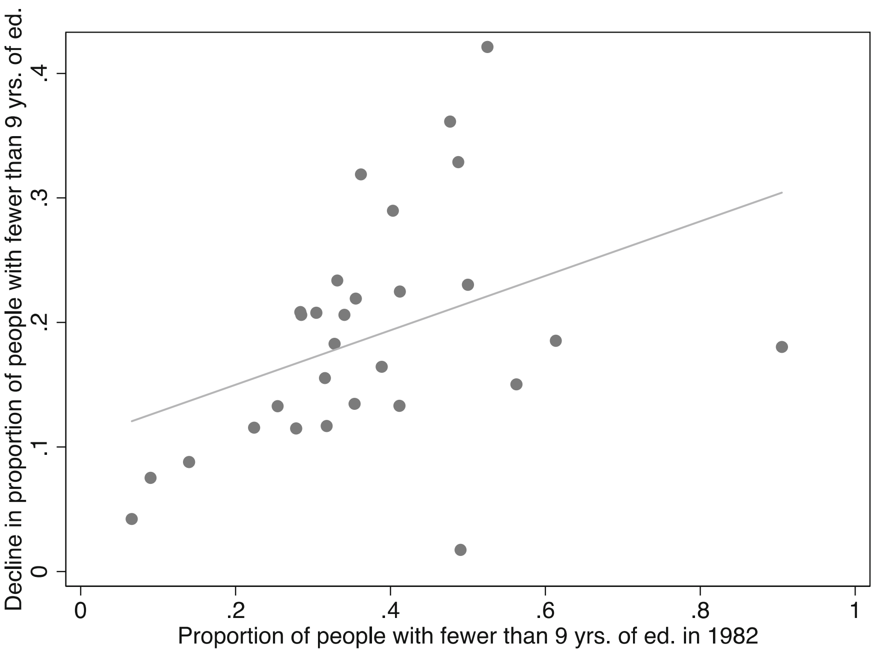
Decline in proportion of people with fewer than 9 yrs. of ed.
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1
Proportion of people with fewer than 9 yrs. of ed. in 1982

◀ **Fig. 2** Province-level educational levels prior to the CSLs and changes over time. *Note*: Authors' own construction based on the 1% sample of the 1982 and the 2000 Population Census data. **a** Pre-reform proportion of individuals (ages 16–18) with fewer than 9 years of schooling across provinces in China. Exact numbers are listed in Appendix 4. **b** Post-reform decline in the proportion of individuals (ages 16–18) with fewer than 9 years of schooling versus pre-reform proportion of individuals (ages 16–18) with fewer than 9 years of schooling across provinces in China. The *y*-axis displays the decline in the proportion of individuals (ages 16–18) with fewer than 9 years of schooling between 1982 and 2000. The *x*-axis is the pre-reform proportions

replaced by the gender role attitudes directly. The estimated equation is.

$$\mathrm{Attitude}_{\mathrm{ipt}} = \gamma_1 \mathrm{prop}_{\mathrm{p}}^{\mathrm{edu}<9} \times \mathrm{Exp}_{\mathrm{pt}} + \gamma_2 X_{\mathrm{ipt}} + D_{\mathrm{ipt}} + D_{\mathrm{effective\ year}} \times D_{\mathrm{t}} + \sum_j x_{\mathrm{p}}^j \times \mathrm{Exp}_{\mathrm{pt}} + v_{\mathrm{ipt}}, \quad (3)$$

where γ_1 captures the effect of the CSLs on the attitudes toward gender roles. Alternatively, we can use the interaction term of $\mathrm{prop}_{\mathrm{p}}^{\mathrm{edu}<9}$ and $\mathrm{Exp}_{\mathrm{pt}}$ as the instrumental variable for years of schooling and apply an IV estimation.[17]

4.4 Impact of the CSLs on economic behavior

The outcome variables used in this paper are the survey responses to gender attitude questions. However, the positive relationship between schooling and egalitarian gender role attitudes may be due to the fact that better-educated individuals know how to properly respond to the questions and hide their "true" attitudes. In this section, we explore the impact of schooling on the economic behavior of individuals. Specifically, we test whether educated women tend to have a more equal division of household chores between men and women in a household. We also investigate whether education affects the fertility behavior and employment status of women. The estimated equation is.

$$\mathrm{Behavior}_{\mathrm{ipt}} = \theta_1 \mathrm{prop}_{\mathrm{p}}^{\mathrm{edu}<9} \times \mathrm{Exp}_{\mathrm{pt}} + \theta_2 X_{\mathrm{ipt}} + D_{\mathrm{ipt}} + D_{\mathrm{effective\ year}} \times D_{\mathrm{t}} + \sum_j x_{\mathrm{p}}^j \times \mathrm{Exp}_{\mathrm{pt}} + \vartheta_{\mathrm{ipt}}, \quad (4)$$

where $\mathrm{Behavior}_{\mathrm{ipt}}$ represents women's involvement in housework, whether a woman has given birth to a second child given the first birth is a girl, and the employment status of a woman, respectively.

5 Estimation results

5.1 Impact of the CSLs on education

5.1.1 Graphical evidence

Our empirical results begin with a simple graph of cohort-specific relationships between junior high school completion dummy and pre-reform educational level ($\mathrm{prop}_{\mathrm{p}}^{\mathrm{edu}<9}$). The shift in the partial correlation between junior high school completion

[17] The estimates from the reduced form regression tend to capture the overall impacts of the reform on gender role attitudes, for example, the increased presence of female teachers after the reform, who serve as a role model for female students, may lead to more progressive gender attitudes. The percentage of female primary school teachers increased from 41.68% in 1982 to 47.69% in 1990.

dummy and pre-reform educational level coincides with our measure of the CSLs reform exposure, which can be seen in this section. We compare changes in junior high school completion by cohort across provinces with distinct pre-reform educational levels in order to assess the contribution of the CSLs to the observed changes.

For a 2-year age group defined by the age when the law was implemented in the province of birth, we regress the indicator of completing 9 years of education on $\mathrm{prop}_{\mathrm{p}}^{\mathrm{edu}<9}$ controlling for individual controls, regional fixed effects, and survey year fixed effects. The estimates of the partial correlation and the 95% confidence intervals are plotted in Fig. 3. To relate these estimates of partial correlations to the measure of the CSLs reform exposure, we overlay each age group's potential eligibility to CSLs as a red split line in the figure. The figure shows that, a person that was not eligible to the CSLs is 45 percentage points more likely to complete junior high school if he/she was born in a province where everyone completed junior high school education in 1982, as compared with his/her counterpart in a province where no one completed junior high school education. For cohorts that are fully eligible to the CSLs, the partial correlation between initial education level and his/her probability of completing junior high school education is around zero. For those who are partially eligible, the partial correlation reduces almost linearly with our measure of reform exposure. The plot implies that the CSLs have eliminated the cross-province variation in compulsory school attendance in 1982 that cannot be explained by individual characteristics and region fixed effects. It also reveals that our measure of the CSLs reform exposure captures the extent to which

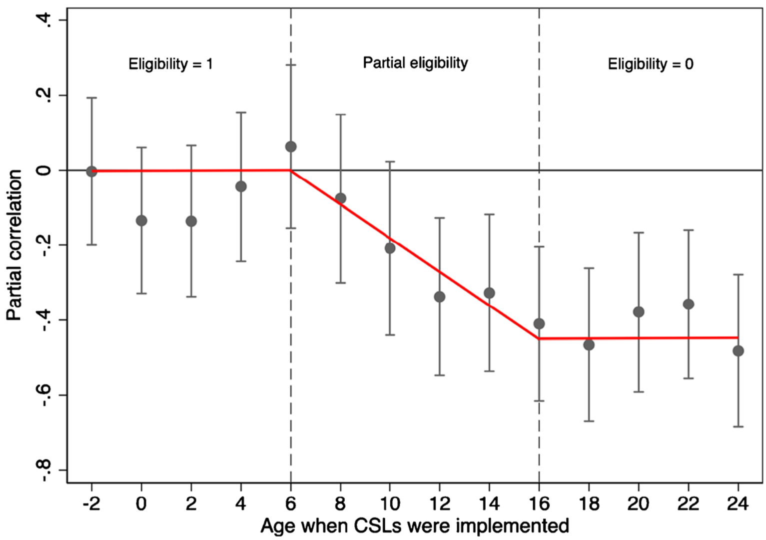

Fig. 3 Partial correlation of initial educational levels and the probability of completing junior high school education, by age group. *Note*: The red lines reassemble the theoretical CSLs eligibility displayed in Appendix 2. The *x*-axis represents an individual's age *t* when CSLs were implemented in the province of birth. The *y*-axis represents the coefficient and confidence interval corresponding to a separate regression for individuals aged *t* and *t*+ 1 when CSLs were implemented, where we regress an indicator of completing junior high school education on the share of junior high school graduates in 1982 in the province of birth, conditional on individual controls, region fixed effects, and survey year fixed effects

a cohort was affected with respect to the probability of completing compulsory schooling.

5.1.2 Baseline regression results

Table 2 reports the OLS estimates of the CSLs' impacts on years of schooling.[18] The dependent variable in columns (1) to (3) is years of schooling. Column (1) presents the baseline results from estimating Eq. (2). The statistically significant and positive coefficient of the interaction term suggests that the law-induced increase in years of schooling is greater in regions with lower pre-reform education. The magnitude of the coefficient implies that the policy-induced increase in years of schooling in low-education regions with a junior high school completion rate of around 0.4 before the CSLs (e.g., Yunnan, Guizhou, and Gansu) would be approximately 1 year higher than the high-education regions with a completion rate of around 0.9 before the CSLs (e.g., Beijing, Tianjin, and Shanghai). The large first-stage F-statistic (approximately 25) indicates that the reform has a strong impact on years of schooling.

Nevertheless, other factors, for example, economic growth or relative school quality improvements, might have caused the differential cohort trends across regions. As shown in Stephens and Yang (2014), when allowing the year of birth effects to vary across regions, the statistically significant causal impact of compulsory education law in previous literature becomes insignificant and, in many instances, wrong-signed. Thus, in column (2), we allow the year of birth effects to vary across the four regions of birth, namely, east coast, central China, northeast China, and western China, by further including region-year of birth fixed effects. The estimated coefficient remains the same after controlling for region-year of birth-fixed effects, suggesting that other factors across different regions should not be the factor driving the results.

To rule out other possible explanations of the differential changes in education across provinces, we further include the interactions between reform exposure and pre-reform characteristics of the province of birth in column (3), which is our preferred specification. The coefficient from our preferred specification indicates that the policy-induced increase in years of schooling in regions with lower prior education is approximately 1.25 years higher than the regions with higher pre-reform education.

The dependent variable in columns (4) to (6) is an indicator variable for the completion of junior high school education. The coefficients of the interaction term are positive and statistically significant. The magnitudes imply that the policy-induced increase in the probability of completing junior high school education in regions with lower prior education is 25–30 percentage points higher than the increase in regions with higher pre-reform education. Because on average 70% of individuals from low-

[18] As a robustness check, we report the results from a difference-in-differences (DD) specification in Appendix 5. Our main specification permits the DD estimate to vary by pre-reform education level, which captures the heterogeneous effects of the DD estimate on children from different socioeconomic backgrounds. Following Black et al. (2005), we restrict the DD analysis to children born into families where the father had less than primary education. These children are most affected by the compulsory education reform. The coefficients of the DD term are significantly positive, suggesting that children from low-educated families tend to comply more with the reform. While not reported, we find no effects on children with educated fathers. This finding is consistent with the triple-difference design, which exploits the fact that people from low-education regions are more likely to be the compliers to the CSLs.

Table 2 Impact of the CSLs on schooling

Variables	(1)	(2)	(3)	(4)	(5)	(6)
	Years of schooling			Completion of junior high education		
Reform exposure × Pr (less than 9 years of education)	2.038*** (0.415)	2.921*** (0.432)	2.523*** (0.491)	0.534*** (0.048)	0.567*** (0.050)	0.504*** (0.057)
Region-birth year FE	No	Yes	Yes	No	Yes	Yes
Reform exposure × province char.	No	No	Yes	No	No	Yes
Observations	21,867	21,867	21,856	21,867	21,867	21,856
R-squared	0.495	0.499	0.498	0.295	0.300	0.300
First-stage *F*-statistic	24.06	45.63	26.41	123.38	127.49	77.28

Note: Each column represents a separate regression from estimating Eq. (2). The dependent variable is years of schooling in columns (1)–(3) and is an indicator variable for the completion of junior high education in columns (4)–(6). All regressions include individual controls and fixed effects for birth year, birth month, survey year, province, province-survey year, and effective year-enrollment year. Individual controls include gender, ethnicity, *hukou* status (urban/rural), age, both parents' years of schooling, and mother's employment status when the respondent was 14. Provincial characteristics include average wage, population density (person/km^2), railway mileage (kilometer), and fiscal expenditure on culture, education, science, and public health in 1985. Robust standard errors in parentheses are clustered at the province-birth year level

*Significant at 10%; **significant at 5%; ***significant at 1%

education regions did not complete junior high school education in 1982, our point estimate suggests that a maximum of about 35% (0.5 × 70%) of full-eligible individuals from low-education regions were induced to complete at least junior high school education.

5.1.3 Heterogeneous effects of the CSLs

The educational level prior to the CSLs remarkably differs across gender and urban/rural groups. Thus, the effect of the CSLs on years of schooling may depend on the gender and *hukou* status of individuals. Table 3 divides the sample by gender and type of *hukou* to examine the heterogeneous impact of the CSLs on education. The results in the first two columns show that the impact of CSLs is significantly larger for women at the 10% level, which are in line with the policy implementation, as women have lower educational attainment before the implementation of the CSLs and thus are more affected by the CSLs. However, the last two columns in Table 3 suggest that the impact of the CSLs for individuals with rural *hukou* is not significantly different from that for individuals with urban *hukou*.

5.1.4 Impact of the CSLs on years of schooling at different educational levels

The CSLs require universal 9 years of compulsory schooling for all provinces. Hence, the reform should have the largest impact on the probability of completing at least 9-year education. In this subsection, we provide information on the educational level at which the program was effective. To this end, we construct a set of indicators for

Table 3 Impact of the CSLs on years of schooling, by gender and *hukou*

Variables	(1)	(2)	(3)	(4)
	Dependent variable: years of schooling			
	Male	Female	Urban	Rural
Reform exposure × Pr (less than 9-year education)	1.705** (0.695)	3.242*** (0.666)	2.644*** (0.826)	2.156*** (0.736)
p value	0.10		0.66	
Observations	10,541	11,315	8964	12,892
R-squared	0.456	0.551	0.338	0.328
First-stage F-statistic	6.03	23.70	10.25	8.59

Note: Each column represents a separate regression from estimating Eq. (2). All regressions include individual controls and fixed effects for birth year, birth month, survey year, province, province-survey year, effective year-enrollment year, region-birth year, and interactions between reform exposure and provincial characteristics. Individual controls include gender, ethnicity, *hukou* status (urban/rural), year of birth, age, both parents' years of schooling and mother's employment status when the respondent was 14. Provincial characteristics include average wage, population density (person/km^2), railway mileage (kilometer), and fiscal expenditure on culture, education, science and public health in 1985. Robust standard errors in parentheses are clustered at the province-birth year level

*Significant at 10%; **significant at 5%; ***significant at 1% (p values are from generalized Hausman specification tests for the difference between estimates for different sub-groups (Hausman 1978))

different levels of education, use these indicators as dependent variables, and conduct the regressions as in Eq. (2).

The coefficients of the interaction term are plotted in Fig. 4 (the 95% confidence interval is plotted by dashed lines). The shape of Fig. 4 indicates at what level the law was effective. The effect is increasing until the 9th year of education and is decreasing afterward. Figure 4 also shows some impacts of the reform on the probability of completing senior high school education. This implies that individuals who completed junior high school education due to the implementation of CSLs may tend to continue to enroll in senior high school. Figure 4 provides additional evidence that the assumption underlying the identification strategy is reasonable as the estimated effect of the program for the levels of education that it did not target is considerably smaller.

5.2 Impacts of education on gender role attitudes

5.2.1 Baseline results

Table 4 reports the impacts of education on gender role attitudes. Panel A in Table 4 reports the OLS estimates of δ_1, implying that years of schooling is positively associated with egalitarian gender role attitudes. Panel B shows the reduced form results using a triple-difference strategy. The significantly positive coefficients of the interaction term in the reduced form regression indicate a positive impact of the CSLs on egalitarian attitudes toward gender roles. Panel C presents the 2SLS estimates, which are of key interest in this paper. The F-tests in the first stage (i.e., weak instrumental variable tests) for the instruments are reported at the bottom of each column. The results

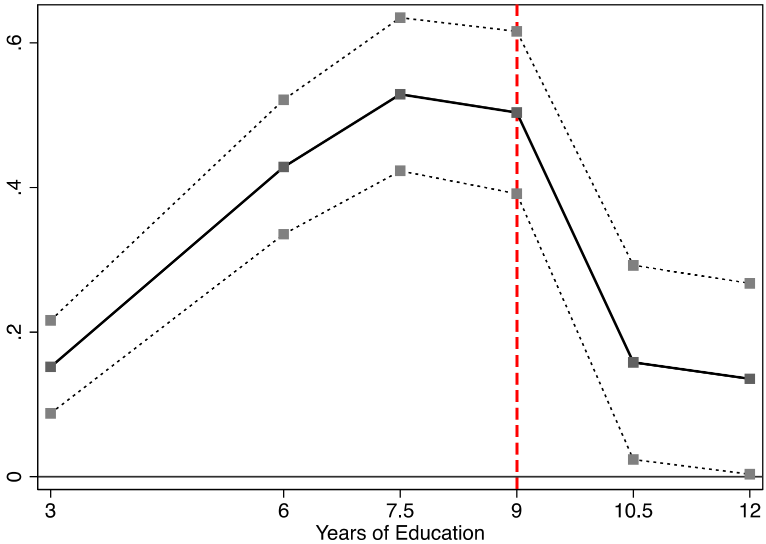

Fig. 4 Impacts of the CSLs on years of schooling at different educational levels. *Note*: The dashed lines indicate the 95% confidence intervals for the regression coefficient of the interaction term

suggest that one additional year of schooling induced by the CSLs increases the probability of the endorsement of egalitarian gender role attitudes by approximately 5–14 percentage points. However, the results in columns (1) and (4) in panel C show that schooling does not have significant effects on the gender role attitudes regarding women's right to a career. The results in column (6) suggest that the impact of increased education on the overall index of gender role attitudes is significantly positive.

OLS tends to overestimate the true effect of schooling on gender role attitudes if parents with egalitarian gender role attitudes tend to invest more in their children's human capital. Nevertheless, there are three potential reasons why IV may exceed OLS. First, the classical measurement error in education may cause OLS estimates to be biased toward zero. However, as the OLS estimates are statistically different from zero in this paper, our results are less likely to be driven by measurement error. Second, the IV estimates reflect the effect of education on the compliers, that is, individuals who respond to changes in schooling laws (i.e., less educated individuals constrained from dropping out earlier). As such, IV estimates measure changes in gender role attitudes associated with lower levels of education (typically junior high school), which may be larger than the average treatment effect. Third, as documented by Lochner and Moretti (2004), when the instrumental variable varies at some aggregate level, for example, at province-birth cohort level in our case, IV estimates capture the combined effect of own education on outcomes as well as the effect of average cohort education on outcomes. That is, IV estimates the sum of the individual effect and the spillover effect, whereas OLS only estimates the individual effect of education.

Table 4 Impact of education on gender-role attitudes

Variables	(1) Career[a] (Egalitarian attitude = 1)	(2) Competency[b]	(3) Marriage[c]	(4) Dismissal[d]	(5) Housework[e]	(6) Overall index[f]
Panel A. OLS estimation						
Years of Schooling	0.021*** (0.001)	0.024*** (0.001)	0.014*** (0.001)	0.014*** (0.001)	0.003*** (0.001)	0.056*** (0.002)
Observations	22,321	22,295	22,272	22,202	22,304	22,109
R-squared	0.102	0.092	0.062	0.092	0.040	0.141
Panel B. Reduced form results						
Reform exposure × Pr (less than 9-year education)	0.075 (0.075)	0.244*** (0.084)	0.346*** (0.080)	0.007 (0.068)	0.118** (0.057)	0.591*** (0.16)
	[0.59]	[0.05]	[0.01]	[0.90]	[0.22]	[0.01]
Observations	21,996	21,970	21,951	21,877	21,979	21,787
R-squared	0.081	0.072	0.056	0.083	0.040	0.113
Panel C. 2SLS estimation						
Years of schooling	0.037 (0.030)	0.101*** (0.039)	0.143*** (0.042)	0.005 (0.027)	0.046* (0.024)	0.254*** (0.079)
Observations	21,829	21,803	21,784	21,712	21,812	21,622
R-squared	0.089			0.089		
First-stage F-statistic	26.07	25.80	26.11	25.61	26.29	25.30

Note: Each cell represents a separate regression. The dependent variable is an indicator of egalitarian gender role attitudes. Romano-Wolf step-down adjusted p values correcting for multiple hypothesis testing are in brackets

*Significant at 10%; **significant at 5%; ***significant at 1%

[a] A dummy variable, which equals to one if the respondent "totally disagrees" or "somewhat disagrees" with the statement: "Men should focus on career, whereas women should focus on family"

[b] A dummy variable, which equals to one if the respondent "totally disagrees" or "somewhat disagrees" with the statement: "Men are naturally more competent than women"

[c] An indicator, equaling to one if the respondent "totally disagrees" or "somewhat disagrees" with the statement: "Getting married with a good man is more important than having a good job"

[d] A dummy variable, which equals to one if the respondent "totally disagrees" or "somewhat disagrees" with the statement: "Female workers should be dismissed first during a recession"

[e] A dummy variable, which equals to one if the answer to the statement: "Husbands and wives should equally share housework" is "neutral," "somewhat agree," or "totally agree"

[f] The standardized sum of five binary items. Robust standard errors in parentheses are clustered at the province-birth year level

5.2.2 Validity tests

The key identification assumption for the triple-difference strategy is that individuals in provinces with similar pre-reform educational levels but different effective dates of the CSLs would have had the same cohort trends in gender role attitudes in the absence of the law. We test this common trend assumption using two placebo tests. In the first placebo test, we restrict the sample to those cohorts who were 17–30 years old when the law took effect, i.e., the cohorts 2–15 years older than the first affected cohort. We re-estimate the triple-difference model as Eq. (3) over the pre-treatment cohorts, but with the assumption that the treatment took effect 5, 6, and 7 years earlier than the actual timing of the CSLs, respectively. The "placebo reform exposure" is constructed based on the assumed timing of the reform in each province and the birth date of each individual. If our empirical specification is valid, the estimates of the placebo treatment variable should be small and statistically insignificant. Table 5 reports the results of placebo tests for the ineligible sample. The coefficients on the interaction between reform exposure and pre-reform education are small and statistically insignificant in the placebo tests, justifying the application of the triple-difference strategy.

The second placebo test is a permutation test over the baseline sample. Specifically, we divide the provinces into groups by the effective year of the CSLs. We randomly assign each province a pre-reform proportion of the population (aged 16–18) with fewer than 9-year schooling from provinces in the same group and construct the "placebo" treatment variable based on the simulated proportion and actual reform exposure. We estimate the placebo treatment effect on the overall index of gender role attitudes for each assignment. To approximate the permutation distribution, we conduct the random assignment 2000 times. The distribution of the placebo estimates is displayed in Fig. 5. The p value of the permutation test is the proportion of placebo estimates greater than or equal to in absolute term the estimated effect of the actual reform from the baseline analysis. The p value is smaller than 0.01, implying that given the same reform exposure, individuals from provinces with lower prior education experienced a significant increase in the endorsement of egalitarian gender role attitudes than those from high-education provinces.

5.3 Robustness checks

In this section, we conduct two sets of robustness checks, namely, more flexible specifications allowing for nonlinear effects of reform exposure, and the inclusion of potential confounding factors.

5.3.1 Nonlinear effects

Our baseline estimation specification applies a linear-in-means model, capturing the linear effects of CSLs exposure. Here, we consider the possibility that the effects of CSLs exposure could be nonlinear and apply more flexible specifications in Appendix 6. First, we add the square of exposure in the regression in column (1). The quadratic term is statistically insignificant, showing no evidence of nonlinearity.

Additionally, we apply categorical dummies rather than a continuous measure in column (2) of Appendix 6. Specifically, we modify column (6) in Table 4

Table 5 Placebo tests for impacts of the CSL

Variables	(1) Career[a]	(2) Competency[b]	(3) Marriage[c]	(4) Dismissal[d]	(5) Housework[e]
Panel A. Suppose CSLs 5 years before					
Reform exposure × Pr (less than 9-year education)	− 0.071 (0.26)	− 0.210 (0.27)	− 0.149 (0.30)	0.124 (0.26)	− 0.053 (0.24)
Observations	10,059	10,045	10,036	9988	10,047
R-squared	0.080	0.077	0.055	0.074	0.046
First-stage *F*-statistic	0.07	0.61	0.24	0.23	0.05
Panel B. Suppose CSLs 6 years before					
Reform exposure × Pr (less than 9-year education)	− 0.010 (0.21)	− 0.048 (0.22)	− 0.131 (0.24)	0.146 (0.21)	− 0.099 (0.19)
Observations	10,059	10,045	10,036	9988	10,047
R-squared	0.080	0.077	0.055	0.074	0.046
First-stage *F*-statistic	0.00	0.05	0.29	0.51	0.28
Suppose CSLs 7 years before					
Reform exposure × Pr (less than 9-year education)	0.038 (0.17)	− 0.003 (0.19)	− 0.160 (0.20)	0.171 (0.17)	− 0.117 (0.15)
Observations	10,059	10,045	10,036	9988	10,047
R-squared	0.080	0.077	0.055	0.074	0.046
First-stage *F*-statistic	0.05	0.00	0.63	0.97	0.60

Note: Each column represents a separate regression from estimating Eq. (3) using the pre-treatment cohorts aged 17–30 years old when the CSLs took effect. The dependent variable is an indicator of egalitarian gender role attitudes. All regressions include individual controls and fixed effects for birth year, birth month, survey year, province, province-survey year, effective year-enrollment year, region-birth year and interactions between reform exposure and provincial characteristics. Individual controls include gender, ethnicity, *hukou* status (urban/rural), year of birth, age, both parents' years of schooling and mother's employment status when the respondent was 14. Provincial characteristics include average wage, population density (person/km^2), railway mileage (kilometer), and fiscal expenditure on culture, education, science, and public health in 1985. Robust standard errors in parentheses are clustered at the province-birth year level

*Significant at 10%; **significant at 5%; ***significant at 1%

[a] A dummy variable, which equals to one if the respondent "totally disagrees" or "somewhat disagrees" with the statement: "Men should focus on career, whereas women should focus on family"

[b] A dummy variable, which equals to one if the respondent "totally disagrees" or "somewhat disagrees" with the statement: "Men are naturally more competent than women"

[c] An indicator, equaling to one if the respondent "totally disagrees" or "somewhat disagrees" with the statement: "Getting married with a good man is more important than having a good job"

[d] A dummy variable, which equals to one if the respondent "totally disagrees" or "somewhat disagrees" with the statement: "Female workers should be dismissed first during a recession"

[e] A dummy variable, which equals to one if the answer to the statement: "Husbands and wives should equally share housework" is "neutral," "somewhat agree," or "totally agree"

by including interactions between pre-reform education and dummy variables for three categories of stages when the law was implemented: (1) junior high school (individuals aged between 12 and 15 when the law took effect); (2)

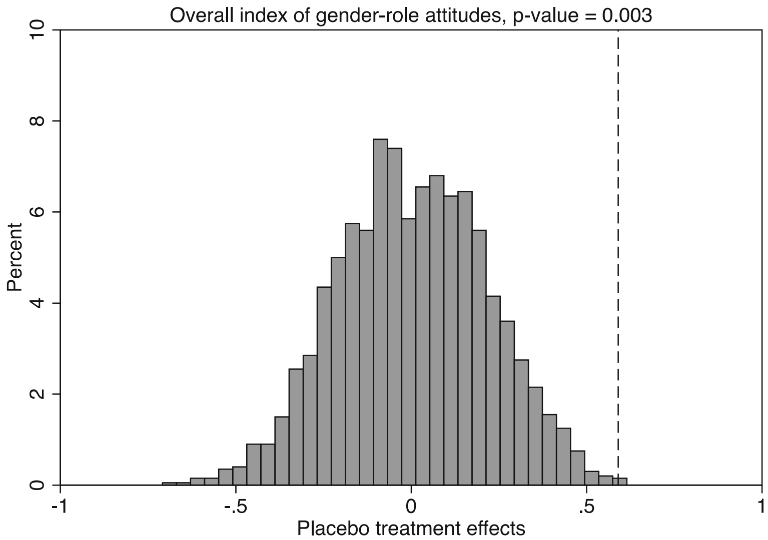

Fig. 5 Estimated coefficients from the permutation placebo tests. *Note*: We randomly assign the pre-reform proportion of people (ages 16–18) with fewer than 9 years of schooling to provinces and construct the "placebo" treatment variable based on the simulated proportion and actual reform exposure, using the baseline sample. The dash line indicates the estimated treatment effect from the baseline analysis. The histogram displays the distribution of the placebo estimates from 2000 random assignments. The *p* value of the permutation placebo test is the proportion of placebo estimates greater than or equal to in absolute value of the estimated effect of the actual reform from the baseline analysis

primary school (individuals aged between 6 and 12 when the law took effect); (3) pre-school (individuals aged below 6 when the law took effect). The omitted category is individuals who are non-eligible for the CSLs. The results shown in column (2) imply that the impact of the CSLs declines with the age of individuals when the law took effect. In addition, the CSLs have the largest impact on the individuals who had not enrolled in primary school when the law took effect. The results in Appendix 6 render support to the baseline linear specification. Moreover, the results also suggest that the baseline results in this paper do not rely on the linear-function assumption.

5.3.2 Confounding factors

The triple-difference strategy assumes that no other shocks that might have affected the outcomes more in areas with lower pre-reform education occurred at the same time as the CSLs. However, China has experienced a series of reforms in the 1980s. Our estimates of CSLs effects will be biased if the roll-out of the CSLs coincides with the timing of other reforms that have heterogeneous impacts varying with local educational levels. To test for the assumption, we re-estimate our IV model while accounting for the impacts of concurrent shocks that may confound our results.

One-child policy The One-Child Policy in China, initiated in 1979, affects the size and gender composition of siblings, which may further influence individuals' educational attainment and gender role attitudes.[19] The enforcement of the One-Child Policy varies across regions and years. If the enforcement coincides with the timing of the CSLs, the omission of the One-Child Policy may bias our results. We measure the strength of the enforcement of the One-Child Policy by using the average monetary penalty rate for one unauthorized birth in the provincial-year panel from 1979 to 2000 from Ebenstein (2010).[20] The fine rates are formulated in years of household income (Ebenstein 2010; Huang et al. 2016; Huang and Zhou 2015). Because a pregnancy usually lasts for 9 months, parents' decision to have a child, if any, should be made close to a year in advance. Therefore, we match the CGSS data with the policy fine (at the provincial level) 1 year before the birth year of individuals. To capture the differential effects of fine rates in urban and rural areas, we add fine rates interacted with rural *hukou* dummy in the regression. The results in panel A of Table 6 demonstrate that after controlling for the enforcement of One-Child Policy, our main results hold. Interestingly, we find that rural individuals born in an environment with stricter restrictions on family planning hold more egalitarian gender role attitudes.

Migration opportunities Since the late 1980s, the *hukou* system has been relaxed, and people in rural areas have been allowed to migrate to urban areas. A large scale of rural-urban migration has occurred in China since the first wave of migration in 1989. The opportunity to migrate may affect individuals' school enrollment (De Brauw and Giles 2017) as well as gender role attitudes. In this study, we use two means to control for migration opportunities for cohorts in different regions. First, the development of urban private and informal sectors in coastal areas boosts migration opportunities for rural residents. In the regression, we have controlled for the region-cohort fixed effects to allow for differential birth year effects across different regions. Second, the migration rate before the relaxation of the *hukou* system, which reflects the initial migration networks, may affect the migration decision of individuals in the late 1980s. We thus further control for migration opportunities by including the interactions between the pre-reform migration rate in the province of birth and birth cohort dummies. Specifically, the migration rate is defined as the average ratio of the number of individuals not living in the household for more than 1 year to the number of the labor force in a household, constructed based on the 1% sample of the 1982 census data. The results reported in panel B of Table 6 suggest that our main results are not confounded by migration.

[19] The theory of a quantity-quality trade-off suggests that the educational attainment of individuals decreases with the size of siblings (Becker 1960). Moreover, the gender composition of siblings has significant impacts on education (Lei et al. 2017). The gender composition of siblings may also affect individuals' gender role attitudes, as in a household with both boys and girls, parents may treat boys and girls differently, whereas in a household with an only child, girls may not feel that they are treated differently, though their parents still hold traditional gender role attitudes.

[20] The provincial average fine rates depended upon provincial regulations, ethnic composition, and the share of parents with urban *hukou* (Ebenstein 2010).

Table 6 Robustness checks—confounding factors

Variables	(1) Career[a]	(2) Competency[b]	(3) Marriage[c]	(4) Dismissal[d]	(5) Housework[e]	(6) Overall index[f]
Panel A. 1-child policy						
Years of schooling	0.021 (0.032)	0.094** (0.041)	0.140*** (0.044)	− 0.015 (0.030)	0.043 (0.026)	0.217*** (0.081)
Fine rate	− 0.021 (0.024)	− 0.015 (0.028)	0.009 (0.031)	− 0.030 (0.019)	0.008 (0.018)	− 0.036 (0.054)
Fine rate × rural *hukou*	0.053*** (0.017)	0.027 (0.021)	0.004 (0.023)	0.068*** (0.016)	0.006 (0.013)	0.119*** (0.041)
Observations	21,829	21,803	21,784	21,712	21,812	21,622
R-squared	0.102			0.054		
Panel B. Migration opportunities						
Years of schooling	0.046 (0.030)	0.107*** (0.039)	0.151*** (0.042)	0.004 (0.028)	0.053** (0.024)	0.275*** (0.081)
Migration rate in 1982 × birth year dummy	Yes	Yes	Yes	Yes	Yes	Yes
Observations	21,829	21,803	21,784	21,712	21,812	21,622
R-squared	0.073			0.089		

Note: The dependent variable is an indicator of egalitarian gender role attitudes. The fine rates are formulated in years of household income. All regressions include individual controls and fixed effects for birth year, birth month, survey year, province, province-survey year, effective year-enrollment year, region-birth year, and interactions between reform exposure and provincial characteristics. Robust standard errors in brackets are clustered at the province-birth year level

*Significant at 10%; **significant at 5%; ***significant at 1%

[a] A dummy variable, which equals to one if the respondent “totally disagrees” or “somewhat disagrees” with the statement: “Men should focus on career, whereas women should focus on family”

[b] A dummy variable, which equals to one if the respondent “totally disagrees” or “somewhat disagrees” with the statement: “Men are naturally more competent than women”

[c] An indicator, equaling to one if the respondent “totally disagrees” or “somewhat disagrees” with the statement: “Getting married with a good man is more important than having a good job”

[d] A dummy variable, which equals to one if the respondent “totally disagrees” or “somewhat disagrees” with the statement: “Female workers should be dismissed first during a recession”

[e] *A* dummy variable, which equals to one if the answer to the statement: “Husbands and wives should equally share housework” is “neutral”, “somewhat agree”, or “totally agree”

[f] The standardized sum of five binary items

5.3.3 Narrower bandwidths

In the main analyses, we restrict the sample to individuals born between 1960 and 1990. As a robustness check, we restrict the sample to narrower ranges of observations around the first affected cohort and test whether the 2SLS estimates are sensitive to narrower ranges of cohorts. In Appendix 7, the dependent variable is the overall index of gender role attitudes. We use a 14-year window (column 1), a 12-year window

Table 7 Impact of education on gender-role attitudes, by gender and *Hukou*

Variables	(1) Career[a]	(2) Competency[b]	(3) Marriage[c]	(4) Dismissal[d]	(5) Housework[e]	(6) Overall index[f]
Panel A. Gender						
Female						
Years of schooling	0.029 (0.035)	0.096** (0.038)	0.162*** (0.046)	0.048* (0.028)	0.025 (0.022)	0.264*** (0.081)
Observations	11,301	11,286	11,280	11,232	11,290	11,189
Male						
Years of schooling	0.063 (0.062)	0.106 (0.080)	0.111 (0.076)	− 0.066 (0.068)	0.074 (0.060)	0.234 (0.16)
Observation	10,528	10,517	10,504	10,480	10,522	10,433
Panel B. *Hukou*						
Urban *hukou*						
Years of schooling	0.079 (0.051)	0.072 (0.053)	0.136** (0.062)	− 0.048 (0.039)	0.032 (0.033)	0.202** (0.10)
Observations	8959	8946	8949	8941	8949	8907
Rural *hukou*						
Years of schooling	− 0.099 (0.068)	− 0.017 (0.062)	0.145* (0.076)	0.024 (0.061)	0.090* (0.052)	0.112 (0.13)
Observations	12,870	12,857	12,835	12,771	12,863	12,715

Note: The dependent variable is an indicator of egalitarian gender role attitudes. Specifically, *career* is a dummy variable, which equals to one if the answer to the question: “Men should focus on career, whereas women should focus on family” is “totally disagree” or “somewhat disagree.” All regressions include individual controls and fixed effects for birth year, birth month, survey year, province, province-survey year, effective year-enrollment year, region-birth year, and interactions between reform exposure and provincial characteristics. Robust standard errors in parentheses are clustered at the province-birth year level

*Significant at 10%; **significant at 5%; ***significant at 1%

[a] A dummy variable, which equals to one if the answer to the question: “Men are naturally more competent than women” is “totally disagree” or “somewhat disagree”

[b] A dummy variable, equaling to one if the answer to the question: “Getting married with a good man is more important than having a good job” is “totally disagree” or “somewhat disagree”

[c] A dummy variable, which equals to one if the answer to the question: “Female workers should be dismissed first during a recession” is “totally disagree” or “somewhat disagree”

[d] A dummy variable, which equals to one if the answer to the question: “Husbands and wives should equally share housework” is “neutral,” “somewhat agree,” or “totally agree”

(column 2), and a 10-year window (column 3) around the first affected cohort. The results are very similar to our main results.

5.4 Heterogeneous effects of schooling

Considering that the effect of schooling may be different for men and women, we conduct 2SLS estimation for men and women, respectively, and report the results in panel A of Table 7. We find that education has a stronger effect on women's gender-related attitudes than on men's, which is consistent with the finding of Kane (1995).[21] The results suggest that men's attitudes have been slower to change than women's. Given that men hold more traditional gender role attitudes in general in China, our results imply that the gap between men's and women's attitudes has widened over time, which is in line with the US evidence (Brewster and Padavic 2000).

Additionally, as people in rural China have more traditional attitudes toward gender roles, education's liberalizing effect may be weaker in rural areas. To examine whether education's liberalizing effect differs across regions, we also conduct the estimation for individuals with urban and rural *hukou*, respectively. The results, reported in panel B of Table 7, show that education's impact on overall gender role attitudes is more pronounced for individuals with urban *hukou*, which is mainly driven by its impact on gender role attitudes toward marriage. However, education has a stronger effect on attitudes regarding the division of household chores for individuals with rural *hukou*.

5.5 Impacts of education on economic behavior

In this section, we explore the impacts of schooling on the economic behavior of individuals. The analysis of behavior can alleviate the concern of reporting error in self-reported gender role attitudes, and also provide evidence on the channels through which education shapes attitudes.[22] Specifically, we test whether educated women, who hold more egalitarian views toward gender roles, tend to have a more equal division of household chores. We also investigate whether education affects fertility decisions and employment status of women.

We first examine the impact of education on the division of housework. As the CGSS does not contain the information regarding time use, we use the data from the China Family Panel Studies (CFPS) survey, which was launched in 2010 by the Institute of Social Science Survey (ISSS) of Peking University in China. Particularly, CFPS 2010 has a time use module, which collects information about how much time individuals spent in various activities. We restrict the sample to couples of which the wife is aged between 20 and 50 years old from the CFPS 2010. In the time use survey, each respondent was asked to report how many hours per day on average he or she spent on housework on weekdays and weekends, respectively. In our sample, husbands on average spend 0.90 h/day on household chores on workdays, whereas wives spend 143.3% more time (2.19 h) on housework. Although 86% of individuals agree that

[21] Kane (1995) finds that education has a stronger effect on women's gender-related attitudes than on men's.
[22] In Sect. 5.6, we discuss the potential mechanisms underlying education's liberalizing effects on gender role attitudes.

husbands and wives should equally share household chores, the gender gaps in household work in China are substantial.

In Appendix 8, we report the impacts of the CSLs on the division of household chores between husbands and wives using a reduced form regression specification. The outcome variables in the first two columns are the number of hours per day married women spent on household chores on weekdays and weekends, respectively. The dependent variable in the last two columns are the relative division of household chores between husbands and wives, that is, the wife's share of the total time spent on household chores. The results in Appendix 8 suggest that the educational attainment of women does not significantly affect neither the number of hours per day married women spent on household chores nor the relative division of household chores between husbands and wives.

We also examine the impacts of education on the decision to have a second child given the first child is a girl and the employment status of women. In China, women with traditional gender role attitudes have a very strong desire to have at least one son. They are more likely to have a second birth if the first birth was a girl. Hence, the fertility decisions of women to some extent reflect their son preference. As the CGSS does not contain the full birth history of women, we examine the impact of education on fertility behavior of women using the 20% sample of the 2005 1% National Population Sample Survey of China (hereafter the 2005 Mini-Census). We restrict the analysis to urban married women aged 20–50 whose first birth is a girl. The results in column (1) of Appendix 9 suggest that urban married women with higher educational attainment have a lower probability of giving birth to the second child given the first birth is a girl. To investigate the impact of education on the employment status of women, we use the sample of women aged 20 to 50 from the 2005 Mini-Census. As reported in column (2) of Appendix 9, education significantly increases the probability of being employed for women. Extra years of schooling induced by the CSLs change women's fertility behavior and market activity.

Note that the changes in individuals' behaviors may not be solely due to the changes in gender role attitudes induced by the CSLs, as years of schooling may also, for example, increase women's earning potential. The increased earning potential may raise the likelihood of being employed or reduce fertility through possible opportunity cost or income effects. This is one of the limitations of this study.

5.6 Mechanisms

Our main results suggest that receiving more education changes individual gender role attitudes, with a stronger effect on females and urban residents. In this section, we discuss the possible mechanisms through which education shapes individuals' gender identity.

5.6.1 Content of education

Education is an institution fostering the development of "modern" world values. Recent studies have examined the impacts of the content of education on students' attitudes. For example, Cantoni et al. (2017) study the effect of educational content on students' beliefs and attitudes and show that studying the new curriculum has led to more positive views of China's governance,

changed views on democracy, and increased skepticism toward free markets. Dhar et al. (2018) demonstrate that school-based intervention that engaged adolescents in classroom discussions about gender equality, made gender attitudes more progressive and produced more gender-equal behavior.

The curriculum used in primary and junior high school may help to prevent a stereotypical view of the roles of men and women in society. Hence, the content of education may shape individuals' gender role attitudes (Lindsey 2005). However, as we do not have detailed information on the curriculum used in primary and junior high schools in China, we could not explicitly test this mechanism in the paper.

5.6.2 Economic value of women

Several studies have shown that the increased economic value of women reduces gender inequality and leads to more gender-equal preferences. For instance, Qian (2008) finds that an increase in the economic value of adult female labor improves survival rates for girls. Jensen (2012) and Heath and Mushfiq Mobarak (2015) demonstrate that an exogenous increase in women's labor force opportunities leads to increased human capital investments for girls and delayed marriage and childbearing for women. Increased employment opportunities also reduce fertility and increase aspirations for a career. In a recent working paper, Xue (2018) finds that the cotton revolution that led to high productivity for women in the phase of proto-industrialization of China have changed gender role attitudes and led to more gender-equal outcomes.

More education, induced by the compulsory education reform, increases the economic value of women in society, which improves women's socioeconomic status and may further foster more egalitarian gender role attitudes. We show in Appendix 9 the positive reduced-form effects of the reform on female labor force participation, providing suggestive evidence that increased labor market participation may be a channel through which education leads to more liberalized views among females. Moreover, we find a stronger effect on females' gender role attitudes than that on males', suggesting the existence of such a channel, likely the increased labor supply of women, that affects females more than males. However, the potential simultaneity between increased labor supply and liberalized gender role attitudes requires further research to pin down the causal effects of labor market participation on gender role attitudes among women.

5.6.3 Media use

People are constantly exposed to various forms of media, which play a large role in forming social norms. A number of studies have demonstrated that information provided by media can influence a wide range of attitudes and behavior. For example, Gerber et al. (2009) find that exposure to newspapers influences voting behavior and turnout behavior. Jensen and Oster (2009) demonstrate that the introduction of cable television increases women's status in rural India. La Ferrara et al. (2012) show a negative effect of the exposure to soap operas in Brazil on fertility.

Educated individuals tend to have easier access to media. Appendix 10 shows the impacts of the CSLs on media use, including newspaper, magazine, broadcast, TV, and Internet. The former three are traditional sources of information. We find that extra

years of schooling induced by the 1986 Compulsory Education Law significantly increases the frequency of newspaper and Internet use. The new information about the outside world and other ways of life provided by various types of media may lead to more egalitarian gender attitudes.

6 Conclusion

Gender inequality favoring males in educational opportunity, access to health care, personal autonomy, household decision making, labor force participation, wage, and promotion are widely observed across countries. Many of these gender gaps persist and the economic development alone is insufficient to ensure significant progress in narrowing the gaps (Duflo 2012). Recent studies have emphasized the importance of cultural norms in perpetuating gender inequality (Alesina et al. 2013; Giuliano 2017). It is of great interest to understand the formation of gender norms. This paper examines whether education plays an important role in shaping individuals' attitudes about the appropriate or natural role of women in society by exploiting the exogenous temporal and geographical variation in the enforcement of the 1986 Compulsory Schooling Laws (CSLs) in China. Using the data from the China General Social Survey (CGSS), we find that the introduction of the CSLs has a significantly positive effect on educational achievement. The IV estimates suggest that receiving more education changes individuals' gender role attitudes. Specifically, the extra schooling induced by the compulsory schooling reform leads to more egalitarian gender role attitudes. Education's liberalizing effect is markedly stronger for females and urban residents. In other words, the results suggest that social norms regarding gender roles are more persistent for men and rural residents, who hold more traditional attitudes.

Education shapes individuals' social norms and beliefs and thus influences the participation of women in activities performed outside home, such as market employment. Moreover, women with higher educational attainment have a lower probability to give birth to the second child given the first birth is a girl. However, the education of women does not significantly affect the relative division of household chores between men and women in a household. The results provide suggestive evidence that the education impacts on gender-equal behavior are much weaker than that on attitudes. Finally, we discuss the potential channels through which education shapes individuals' gender role attitudes. The liberalizing effect of education may operate through providing gender-neutral curriculum, increasing the economic value of women, and changing the way people get access to new information outside.

Nevertheless, the effect of education on norms and beliefs in this paper is still a black box. Future research may further explore what elements of schooling, for example, schooling environment (such as the gender of teachers, peer group) or the content of education, are particularly important in shaping individuals' gender role attitudes.

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Compliance with ethical standards

Conflict of interest The authors declare that they have no conflict of interest.

Appendix 1

Fig. 6

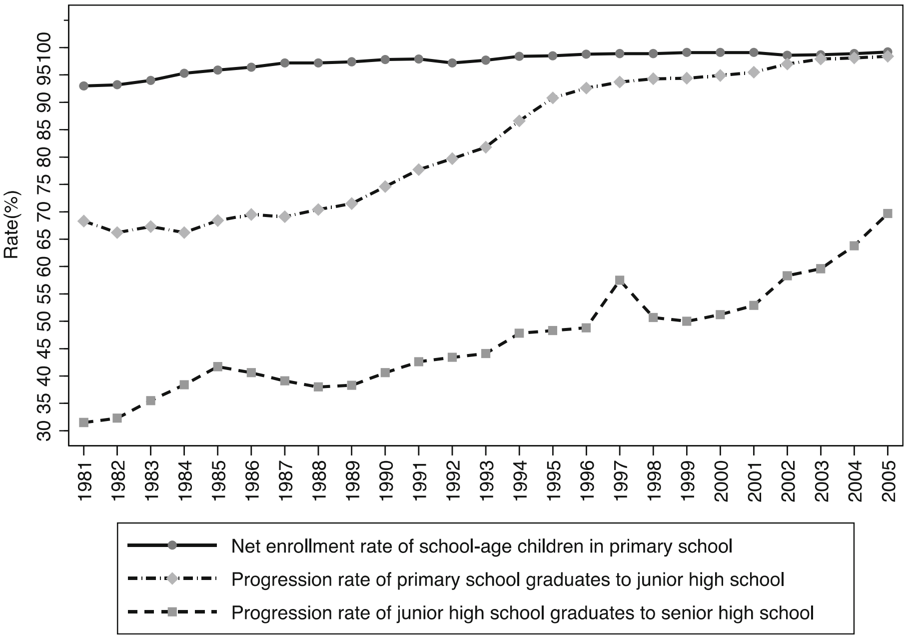

Fig. 6 School enrollment rate and promotion rate in China 1981–2005. *Note:* Time-serious data on primary and junior high school enrollment between 1981 and 2005. *Source*: The China Statistical Yearbook 2006

Appendix 2

Fig. 7

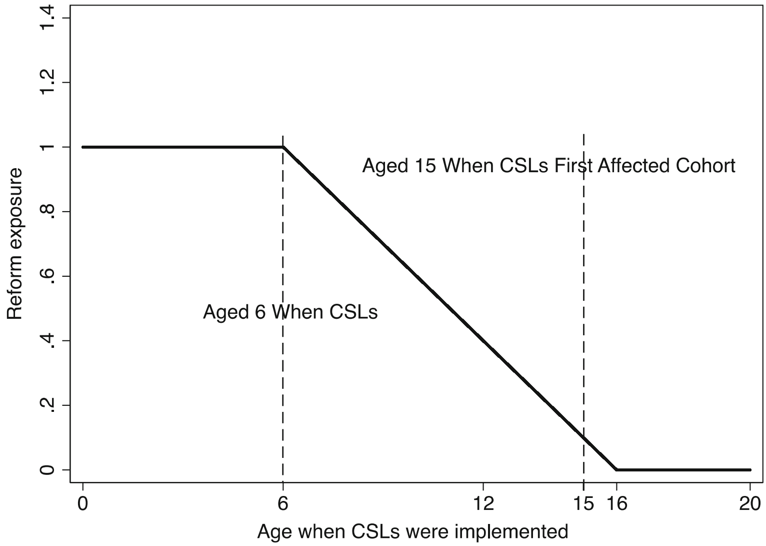

Fig. 7 CSLs eligibility construction. *Note*: The x-axis is the individual age when CSLs were just effective in the local province, and the y-axis is the value for the reform exposure, which equals to 1 if the individual is fully eligible to the CSLs (i.e., aged 6 or below) and equals to zero if the individual is ineligible (i.e., aged 15 or above). A linear function is assumed for the ages in between

Appendix 3

Table 8

Table 8 Summary statistics of main variables

Variables	Mean (1)	Standard deviation (2)	Observations (3)
Panel A. Gender-role attitudes			
(1) Men should focus on career, whereas women should focus on family (dummy)	0.283	0.450	24,973
(2) Men are naturally more competent than women (dummy)	0.455	0.498	24,935
(3) Getting married with a good man is more important than having a good job (dummy)	0.360	0.480	24,896
(4) Female workers should be dismissed first during a recession (dummy)	0.735	0.442	24,789
(5) Husbands and wives should equally share housework (dummy)	0.864	0.343	24,947
Panel B. Regressor of interest			
Years of schooling	9.632	4.371	24,806
Education categories			
No education (dummy)	0.054	0.225	24,993
Primary education (dummy)	0.189	0.392	24,993
Junior high (dummy)	0.338	0.473	24,993
Senior high or above (dummy)	0.420	0.494	24,993
Individual and household characteristics:			
Male (dummy)	0.477	0.499	25,009
Age	39.12	8.666	25,009
Han ethnicity (dummy)	0.908	0.289	24,979
Rural *hukou* (dummy)	0.598	0.490	24,707
Father's years of schooling	6.000	4.490	24,157
Mother's years of schooling	4.148	4.376	24,372
Mother's employment status when the respondent was 14 (dummy)	0.854	0.353	24,221

Note: Based on 2010, 2012, 2013, and 2015 waves of the CGSS. Sample consists of individuals born between 1960 and 1990

Appendix 4

Table 9

Table 9 Effective date of the Compulsory Schooling Law in different provinces

Province	Law effective date	Affected cohort born from	Proportion of earlier cohorts aged 16–18 with less than 9-year education
Beijing	8 July 1986	1 September 1971	0.066
Tianjin	12 November 1986	1 September 1972	0.140
Hebei	1 July 1986	1 September 1971	0.285
Shanxi	1 July 1986	1 September 1971	0.254
Inner Mongolia	15 September 1988	1 September 1974	0.318
Liaoning	1 July 1986	1 September 1971	0.224
Jilin	20 February 1987	1 September 1972	0.315
Heilongjiang	1 July 1986	1 September 1971	0.278
Shanghai	10 September 1986	1 September 1972	0.090
Jiangsu	9 September 1986	1 September 1972	0.284
Zhejiang	1 September 1985	1 September 1971	0.362
Anhui	1 September 1987	1 September 1973	0.477
Fujian	1 August 1988	1 September 1973	0.525
Jiangxi	1 February 1986	1 September 1971	0.488
Shandong	12 September 1986	1 September 1972	0.331
Henan	1 October 1986	1 September 1972	0.304
Hubei	1 March 1987	1 September 1972	0.341
Hunan	1 September 1991	1 September 1977	0.355
Guangdong	7 October 1986	1 September 1972	0.403
Guangxi	1 September 1991	1 September 1977	0.389
Hainan	16 December 1991	1 September 1977	
Chongqing	1 July 1986	1 September 1971	0.412
Sichuan	1 July 1986	1 September 1971	0.412
Guizhou	1 January 1988	1 September 1973	0.563
Yunnan	29 October 1986	1 September 1972	0.614
Tibet	1 July 1994	1 September 1979	0.905
Shaanxi	1 September 1987	1 September 1973	0.328
Gansu	3 September 1990	1 September 1976	0.500
Qinghai	1 October 1988	1 September 1974	0.491
Ningxia	1 July 1986	1 September 1971	0.412
Xinjiang	28 May 1988	1 September 1973	0.354

Note: Authors' own construction based on various local legal and administrative documents. The proportion of earlier cohorts aged 16–18 with less than 9-year education is constructed based on 1% sample of 1982 census data

Appendix 5

Table 10

Table 10 Impacts of the CSLs on schooling—DD specification

Variables	(1) Years of schooling	(2) Years of schooling
Reform exposure	1.856*** (0.70)	1.484** (0.66)
Region-birth year FE	No	Yes
Observations	6938	6938
R-squared	0.317	0.328

Note: We restrict the analysis to high impact sample, which includes those born into families where the father had less than primary education. All regressions include individual controls and fixed effects for birth year, birth month, survey year, province, and province-survey year. Individual controls include gender, ethnicity, *hukou* status (urban/rural), year of birth, age, both parents' years of schooling and mother's employment status when the respondent was 14. Robust standard errors in parentheses are clustered at the province-birth year level

*Significant at 10%; **significant at 5%; ***significant at 1%

Appendix 6

Table 11

Table 11 Non-linear effect of CSLs exposure

Variables	(1)	(2)
	Overall index of gender role attitudes	
Reform exposure × Pr (less than 9-year education)	−0.250 (0.69)	
Reform exposure squared × Pr (less than 9-year education)	0.860 (0.69)	
Junior high × Pr (less than 9-year education)		0.238 (0.18)
Primary × Pr (less than 9-year education)		0.329* (0.19)
Preschool × Pr (less than 9-year education)		0.607*** (0.17)
Observations	21,787	21,787
R-squared	0.113	0.113

Note: The dependent variable is the overall index of gender role attitudes. In column (1), pre-education is interacted with reform exposure and reform exposure squared. In column (2), pre-education is interacted with three categories of stages when the law was implemented. All regressions include individual controls and fixed effects for birth year, birth month, survey year, province, province-survey year, effective year-enrollment year, region-birth year, and interactions between reform exposure and provincial characteristics. Individual controls include gender, ethnicity, *hukou* status (urban/rural), year of birth, age, both parents' years of schooling and mother's employment status when the respondent was 14. Provincial characteristics include average wage, population density (person/ km^2), railway mileage (kilometer), and fiscal expenditure on culture, education, science, and public health in 1985. Robust standard errors in parentheses are clustered at the province-birth year level

*Significant at 10%; **significant at 5%; ***significant at 1%

Appendix 7

Table 12

Table 12 Robustness checks—narrower bandwidth

	(1)	(2)	(3)
	Dep. var.: overall index of gender role attitude		
Variables	[− 14, 14]	[− 12, 12]	[− 10, 10]
Years of schooling	0.288*** (0.094)	0.278*** (0.10)	0.276** (0.14)
First-stage F-statistic	19.16	14.41	7.17
Observations	19,690	18,132	15,709

Note: The dependent variable is the overall index of gender role attitudes. Robust standard errors in parentheses are clustered at the province-birth year level

*Significant at 10%; **significant at 5%; ***significant at 1%

Appendix 8

Table 13

Table 13 Impact of education on division of household chores

Variables	Wife's hours on household chores		Wife's share of time on household chores	
	(1)	(2)	(3)	(4)
	Workday	Weekend	Workday	Weekend
Reform exposure × Pr (less than 9-year education)	0.568 (0.540)	0.856 (0.528)	0.025 (0.127)	0.102 (0.104)
Observations	5505	5500	4338	4391
R-squared	0.127	0.112	0.096	0.113

Note: Based on sample of married women aged between 20 and 50 from CFPS 2010, which collects information about how much time individuals spent in various activities, including personal life (sleep, meals, personal hygiene, household chores, and taking care of family members); individual work; study; entertainment and social activities; transportation; and other. Each column represents a separate regression with specification (4). The dependent variable in columns (1) and (2) is wife's hours spent on household chores and is wife's share of total hours spent on household chores in a household in columns (3) and (4). All regressions include individual controls and fixed effects for birth year, birth month, survey year, province, province-survey year, effective year-enrollment year, region-birth year, and interactions between reform exposure and provincial characteristics. Robust standard errors in parentheses are clustered at the province-birth year level

*Significant at 10%; **significant at 5%; ***significant at 1%

Appendix 9

Table 14

Table 14 Impact of education on fertility and female labor supply

Variables	(1) Having another child	(2) Employment status of women
Reform exposure × Pr (less than 9-year education)	− 0.449*** (0.077)	0.074** (0.033)
Observations	28,826	302,535
R-squared	0.092	0.073

Note: Data source is 2005 Mini-Census. Each column represents a separate regression from estimating Eq. (4). The dependent variable in columns (1) and (2) is an indicator of whether a woman gives birth to another child given the first birth is a girl and the employment status of women, respectively. Column (1) uses the sample of urban married women aged between 20 and 50 whose first birth is a girl; column (2) is based on the sample of women aged between 20 and 50. All regressions include individual controls and fixed effects for birth year, birth month, survey year, province, province-survey year, effective year-enrollment year, region-birth year, and interactions between reform exposure and provincial characteristics. Robust standard errors in parentheses are clustered at the province-birth year level

*Significant at 10%; **significant at 5%; ***significant at 1%

Appendix 10

Table 15

Table 15 Mechanism: impacts of schooling on media use

Variables	(1) Newspaper	(2) Magazine	(3) Broadcast	(4) TV	(5) Internet
Reform exposure × Pr (less than 9-year education)	0.156** (0.072)	0.034 (0.075)	− 0.003 (0.067)	0.018 (0.049)	0.122* (0.070)
Observations	22,016	22,009	21,985	22,006	22,004
R-squared	0.197	0.157	0.136	0.052	0.429

Note: The dependent variable is the frequency of media use, which equals one if the respondent uses the corresponding media sometimes, often, or very often. All regressions include individual controls and fixed effects for birth year, birth month, survey year, province, province-survey year, effective year-enrollment year, region-birth year, and interactions between reform exposure and provincial characteristics. Individual controls include gender, ethnicity, *hukou* status (urban/rural), year of birth, age, both parents' years of schooling and mother's employment status when the respondent was 14. Provincial characteristics include average wage, population density (person/km^2), railway mileage (kilometer), and fiscal expenditure on culture, education, science, and public health in 1985. Robust standard errors in parentheses are clustered at the province-birth year level

*Significant at 10%; **significant at 5%; ***significant at 1%

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